Gasoline Price Wars: Spatial Dependence Awakens

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Abstract

We build an Asymmetric Spatial Error Correction Model (ASpECM) to investigate the role of spatial dependence at the retail gasoline price adjustment mechanism. We find evidence that the symmetric price pattern is fully reversed when we account for spatial spillover effects, indicating that retail prices adjust more rapidly in an upward than a downward direction. This finding raises the possibility that retailers are more likely to engage in anti-competitive practices which may be ignored when the regulators bypass the role of spatial dependence.

JEL classification**:** L13, C23

Keywords: ASpECM; Spatial dependence; Asymmetric gasoline price adjustment

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1. Introduction

Within the last years there is a plethora of studies examining the existence and the causes of gasoline price asymmetry with controversial results (Borenstein et al., 1997; Galeotti et al., 2003; Deltas, 2008; Greenwood-Nimmo and Shin, 2013, Remer, 2015).

Despite the rich body of literature, existing studies suffer from three major limitations. First, they assume that asymmetric price responsiveness is space invariant. However, this is a rather strong assumption that has to be tested rather than assumed in order to avoid biased results. Second and most importantly, all of the existing studies ignore the role of "*localized competition*" between petrol stations and possible spatial spillover effects which in our model act as a driving force to uncover asymmetric price movements. Specifically, Chamberlin (1948) was the first who resembled the gasoline market as a prototype for what he called '*localized competition*'. Slade (1992) has also pointed out that at the retail level, consumers face transportation (time) costs when switching between gasoline stations. Therefore, the gasoline station's location introduces spatial differentiation into a homogeneous product market (Firgo et al., 2015). Thirdly, much of the empirical literature has worked with monthly/weekly data. These datasets have many drawbacks (i.e., less information, inconsistency) since the use of monthly or weekly averages takes away much price variation. Similarly to Remer (2015), we deal with this issue by using daily data. The rest of the paper is organized as follows. The next section describes the empirical model. Section 3 discusses the empirical findings, while Section 4 concludes the paper.

2. The Asymmetric Spatial Error Correction Model (ASpECM)

Extending Beenstock and Felsenstein (2010), the ASpECM is built around a simple long-run relationship of the form:

$$NRPG_{it} = \gamma_{0i} + \gamma_1 SPG_{it} + \gamma_2 NRPG_{it}^* + \gamma_3 SPG_{it}^* + u_{it} \quad (1)$$

where $NRPG_{it}$ is the natural logarithm of the retail price of regular gasoline net of taxes and duties; SPG_{it} is the natural logarithm of the spot gasoline (wholesale) price, γ_{0i} and u_{it} denote the cross-section fixed effects and the error term, respectively and superscript '*' denotes a spatially lagged variable defined as $x_{it}^* = \sum_{h \neq i}^{N} w_{ih} x_{ht}$ for $x = SPG, NRPG$ where N is the number of spatial units and w_{ih} are row-summed spatial weights with $\sum_i w_{ih} = 1$. The matrix $w_{ih} = 0$ if $i = h$ or if $i \neq h$ and the driving distance between spatial units i and h is more than 8.58 *km*. If the distance is less or equal to 8.58 *km*, then $w_{ih} = 1$. Our distance threshold value is calculated by using the speed of the average driver and the fact that a travel time of more than ten minutes is not a good basis for assessing local competition between gasoline stations (OECD, 2011).[1] From (1), we can easily construct the following ASpECM that takes the following form:

$$\begin{aligned}\Delta NRPG_{it} &= a_i + \sum_{j=0}^{k} \delta_j^+ \Delta SPG_{it-j}^+ + \sum_{j=0}^{l} \delta_j^- \Delta SPG_{it-j}^- + \sum_{j=1}^{p} c_j \Delta NRPG_{it-j} + \sum_{j=0}^{q} \delta_j^{*+} \Delta SPG_{it-j}^{*+} \\ &+ \sum_{j=0}^{m} \delta_j^{*-} \Delta SPG_{it-j}^{*-} + \sum_{j=1}^{s} c_j^* \Delta NRPG_{it-j}^* + \lambda^+ ECM_{it-1}^+ + \lambda^- ECM_{it-1}^- + \lambda^{*+} ECM_{it-1}^{*+} \\ &+ \lambda^{*-} ECM_{it-1}^{*-} + v_{it} \quad (2)\end{aligned}$$

[1] However, we use for robustness check possible other spatial weights (i.e., +/-1% of 8.58 km) with similar results.

where t denotes the time dimension and i identifies the spatial units; a_i denotes the spatial specific effect and v_{it} is the usual random i.i.d. error term; Δ is the first difference operator and the orders *k*, *l*, *p*, *q*, *m*, *s* represent the number of lagged terms. $\Delta y_{it}^{+} = \max\{ y_{it} - y_{it-1}, 0\}$ and $\Delta y_{it}^{-} = \min\{ y_{it} - y_{it-1}, 0\}$ for $y = SPG, ECM$, superscript '*' denotes a spatially lagged variable defined as above for $x = SPG, NRPG, ECM$. Lastly, ECM_{it} denotes the error correction term. The coefficients in (1) and (2) are estimated by a fixed effects estimator[2].

3. Estimation results and discussion

The econometric estimation was based on a panel of seven municipalities in the Hudson County of New Jersey (NJ) covering the period from January 2012 to December 2015 (see Figure 1).[3] The choice of the geographical area is based on the fact that NJ is the second State with the highest number of gasoline stations per square kilometer, pointing out possible spatial dependence.

[Insert Figure 1]

We use a daily dataset of 10.199 observations. All price variables constitute daily averages for each one of the petrol stations located on the seven municipalities and are taken from the Oil Pricing Information Service (OPIS). Figure 2 depicts the relatively close co-movement between the wholesale and retail gasoline prices during the sample period.

[Insert Figure 2]

[2] The presence of a very large T (1457) justifies the Least Squares Dummy Variable Estimator (Judson and Owen, 1999).
[3] These include Bayonne, Kearny, North Bergen, Weehawken, Union City, Secaucus and Jersey City. The rest of the municipalities were not included due to missing data.

We carry out the first part of the empirical analysis by examining the presence of cross-sectional dependence (Pesaran, 2004). The test statistic rejects the null hypothesis (*p*-value = 0.000), indicating that price variables exhibit cross-sectional (spatial) dependence. In the next stage, we employ unit root tests. The results support the presence of a unit root (see Table 1, Panel A). The next step is to check for the existence of long-run relationship. According to Pedroni's (1999) panel cointegration test statistics, it is evident that all four tests reject the null hypothesis of no cointegration in the ECM (non-spatial model) and the ASpECM, respectively (see Panel B).

[Insert Table 1]

Table 2 depicts the empirical findings.[4] Examining the non-spatial model, it is evident that wholesale positive coefficient is larger than its negative counterpart, indicating that the effects of upstream price increases are larger than those of price decreases. The coefficients λ^{+} and λ^{-} indicate the asymmetric adjustment speed. From the reported values, we argue that the speed of adjustment ranges from 3.1%-11.4% per day with negative changes of the error correction term being larger (in absolute value), than the positive ones. This means that if the retail price is 10% above its long-run equilibrium, only 0.3% of the difference between the equilibrium and the current price will be eliminated in the next day. We find no evidence of price asymmetry, since the Wald test cannot reject the null hypothesis, while the reverse holds for the existence of long-run asymmetry. If we try to compare the two models, some interesting remarks emerge. First, in the ASpECM, we find a positive (λ^{*+}) and negative (λ^{*-}) long-run coefficient equal to -0.127 and -0.053 respectively, indicating

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[4] We have also estimated the two models using time dummies to control for seasonal effects. The estimation results are qualitatively similar in both cases.

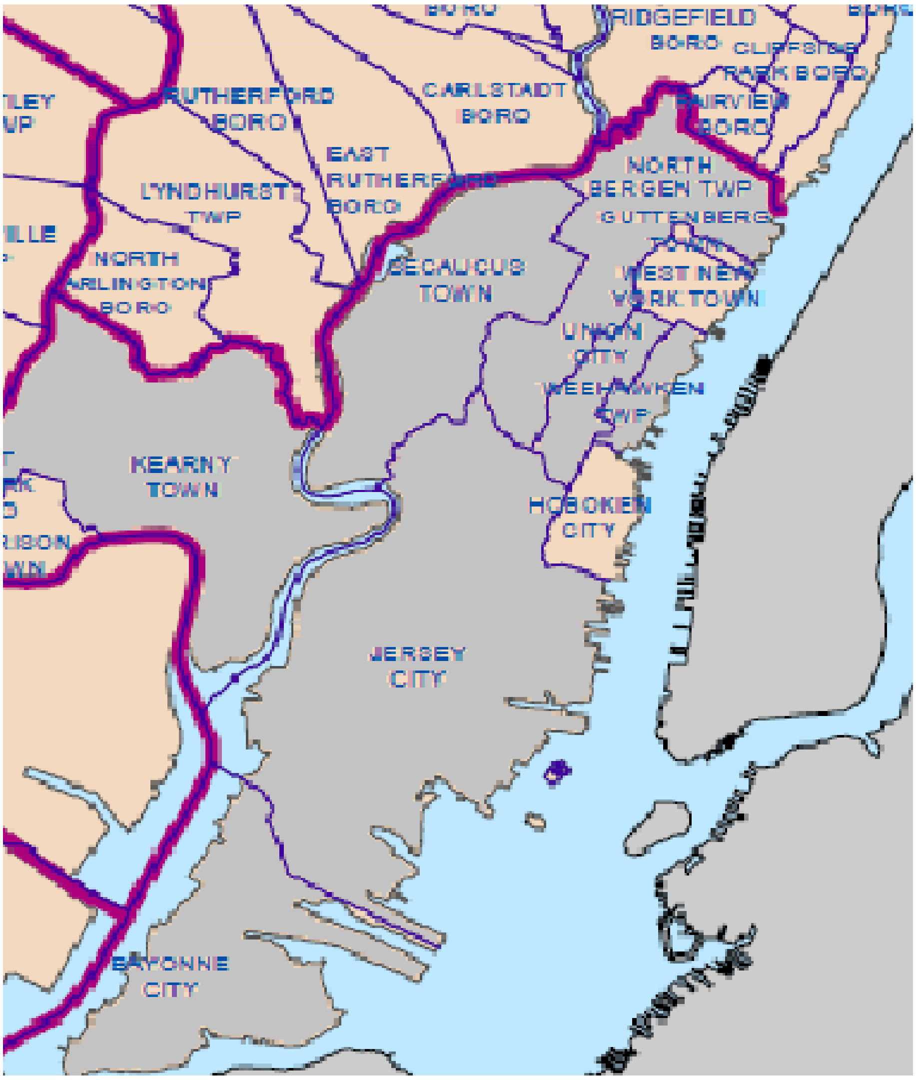
that retailers are driven by the fluctuations in the input price of gasoline. This result reveals a long-run rent-seeking oligopolistic pricing behaviour by the retailers, which in turns is consistent with an asymmetric price adjustment mechanism. This is also confirmed by the rejection of the null hypothesis ($H_0 : \lambda^{*+} = \lambda^{*-}$) giving strong theoretical evidence that spatial asymmetric price adjustment can be attributed to the oligopolistic pricing behavior (Radchenko, 2005). Secondly, the empirical findings of the ASpECM indicate the existence of short-run asymmetric pricing scheme not captured by the ECM since the two Wald tests reject the null. Lastly, the existence of long-run asymmetry in the non-spatial model is strongly biased since the Wald test cannot reject the null in the ASpECM when we do account for spatial dependence (*p*-value = 0.817). Lastly, the diagnostics confirm the absence of serial dependence and the joint statistical significance of the estimated coefficients.

[Insert Table 2]

4. Concluding remarks

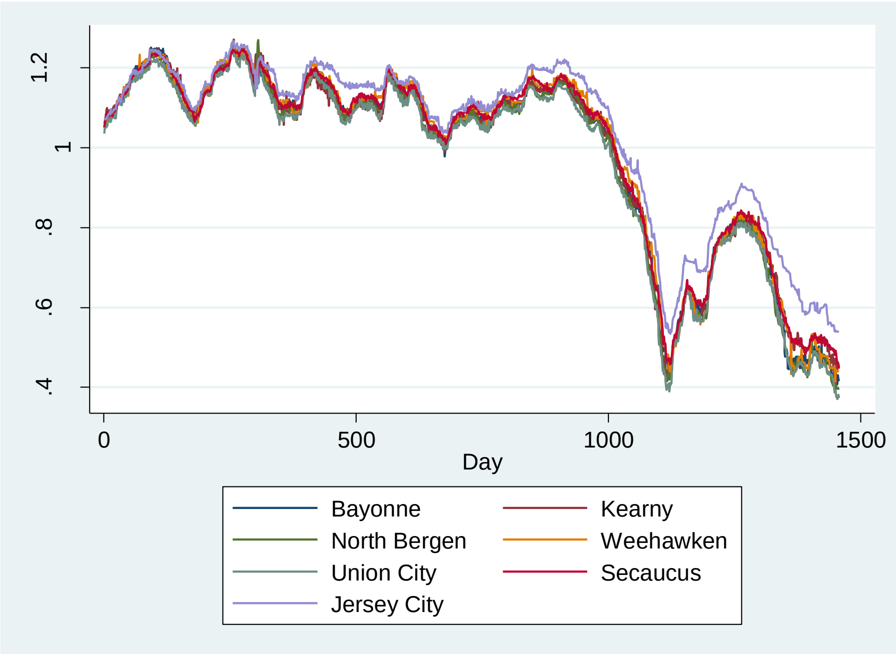
Using the ASpECM framework, we have found strong evidence suggesting the validity of the "*rockets and feathers*" hypothesis. The oligopolistic structure of the local gasoline market along with spatial dependence, trigger the price asymmetric adjustment path. This finding raises serious doubts on the existence of a rent seeking oligopolistic behavior by petrol stations. The difference in the existence of asymmetric pass-through suggests that empirical studies that ignore the role of spatial dependence and local competition may miss an important element of the nature of price adjustment in the retail gasoline industry thus providing the wrong signal to regulators.

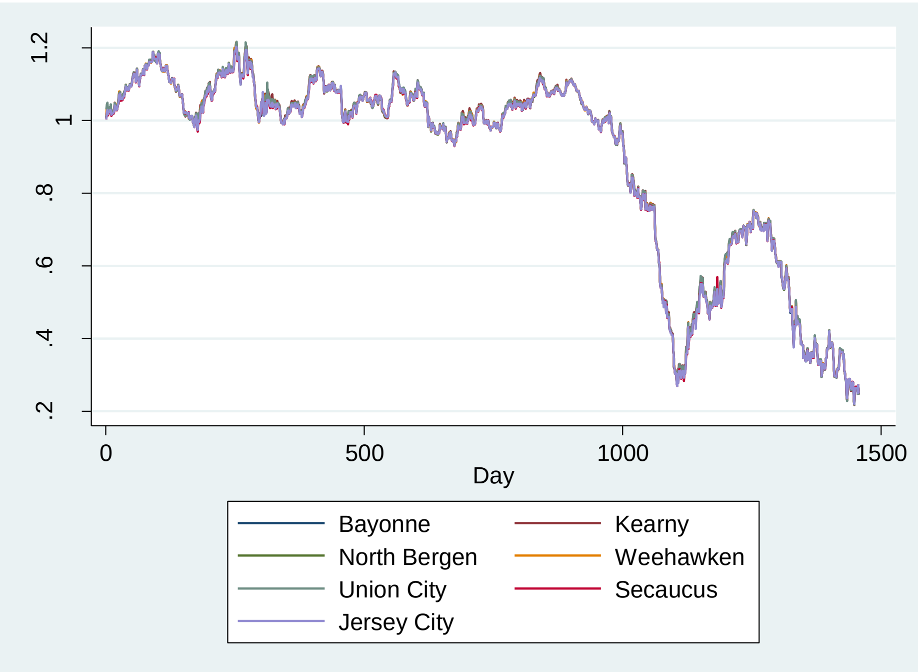
Figure 1: Hudson County municipalities



Source: State of New Jersey Department of State

Figure 2: Price evolution (Jan 2012-Dec 2015)





Source: OPIS Retail Data House

Table 1: Panel unit root and co-integration tests

Panel A: Unit root tests				
Variable	**IPS test**		**Breitung test**	
	W-t-bar	*p*-value	lambda	*p*-value
NRPG	0.113	0.545	0.509	0.695
SPG	2.227	0.987	0.530	0.702
ΔNRPG	-21.930***	0.000	-12.366***	0.000
ΔSPG	-1.2e+02***	0.000	-9.977***	0.000

Panel B: Co-integration tests		
	ECM	**ASpECM**
Statistic		
panel-ρ statistic	-109.6***	-91.98***
panel-t statistic	-32.97***	-34.87***
group-ρ statistic	-120.8***	-118.6***
group-t statistic	-37.72***	-42.76***

Notes: ***, indicate significance at the 1% level. For the Im–Pesaran–Shin (IPS) test, the optimal lag length was chosen by the Bayesian Information Criterion (*BIC*). Breitung's lambda is robust to cross-sectional dependence.

Table 2: Empirical results

Estimated coefficients	**ECM**	**ASpECM**
δ_0^+	0.070*** (0.010)	0.001 (0.019)
δ_0^-	0.063*** (0.009)	0.068*** (0.018)
δ_0^{*+}	-	0.073*** (0.022)
δ_0^{*-}	-	0.010 (0.020)
λ^+	-0.031*** (0.004)	-0.084*** (0.005)
λ^-	-0.114*** (0.004)	-0.086*** (0.006)
λ^{*+}	-	-0.127*** (0.016)
λ^{*-}	-	-0.053*** (0.011)
Symmetry tests		
$H_0: \delta_0^+ = \delta_0^-$	0.19 [0.659]	5.54** [0.019]
$H_0: \delta_0^{*+} = \delta_0^{*-}$	-	3.66* [0.056]
$H_0: \lambda^+ = \lambda^-$	146.82*** [0.000]	0.05 [0.817]
$H_0: \lambda^{*+} = \lambda^{*-}$	-	10.21*** [0.001]
Diagnostics		
Wald's joint test	3440.80*** [0.000]	3192.02*** [0.000]
Autocorrelation of residuals test	0.001 [0.972]	0.050 [0.830]

Notes: ***, ** and * indicate significance at the 1%, 5% and 10% level, respectively. Superscript '*' in δ and λ denotes spatially lagged variables. Standard errors (*p*-values) are reported in parentheses (square brackets). The ECM and the ASpECM are estimated with Feasible Generalized Least Squares (FGLS) with cross-sectional fixed effects, correcting for heteroskedasticity and autocorrelation. The constant term is not statistically significant. The lag length is determined using a recursive procedure based on the value of the *t*-statistic of the last lag for each coefficient. The autocorrelation test is based on Wooldridge (2002).

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